

Regional Integration and the Prices of Imports

An Empirical Investigation

L Alan Winters

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For the first time the effects of regional trading preferences on members and on the rest of the world are investigated by examining their effects on the prices at which trade occurs. The case studies how Spanish accession to the European Community affected the prices of imports from the European Community, Japan and the United States.

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Summary findings

Winters and Chang explore the effects on the terms of trade of regional economic integration. They show why it is an appropriate measure of the welfare effects of such integration, comparing it with the many *ex post* studies that base their conclusions on changes in the import shares of member and nonmember countries.

They demonstrate, by using a simple strategic model, how member countries might gain in their terms of trade, and nonmembers lose, through a lowering of preferential tariffs. Most important, they show that measuring such price effects, though difficult, is feasible.

This is the first *ex post* study of its kind, they believe, and an improvement over previous *ex post* studies on how integration affects the rest of the world. Using finely disaggregated data about Spanish imports of finished manufactures from major OECD trading partners, despite their noisiness, they found a consistent story over all of the country pairs examined.

They find that nonmembers (in this case, Japan and the United States) suffered detectable losses in terms of trade relative to European Community competitors in Spanish import markets for differentiated goods.

An earlier version of this paper — a product of the International Trade Division, International Economics Department — was presented at a North American Economics and Finance Association session during meetings of the Allied Social Science Association, January 1996, in San Francisco. Copies of the paper are available free from the World Bank, 1818 H Street NW, Washington, DC 20433. Please contact Jennifer Ngaine, room N5-056, telephone 202-473-7947, fax 202-522-1159, Internet address jngaine@worldbank.org. June 1997. (39 pages)

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Foreword

As regional trading arrangements (RTAs) have spread, enlarged and deepened over the last decade, they have posed challenges to economists on both intellectual and policy levels. On the former, do RTAs stimulate growth and investment, facilitate technology transfer, shift comparative advantage towards high value-added activities, provide credibility to reform programs, or induce political stability and cooperation? Or do they, on the other hand, divert trade in inefficient directions and undermine the multilateral trading system?

The answer is probably “all of these things, in different proportions according to the particular circumstances of each RTA.” This then poses the policy challenge of how best to manage RTAs in order to get the best balance of benefits and costs. For example, should technical standards be harmonized and, if so, how; do direct or indirect taxes need to be equalized; how should RTAs manage their international trade policies in an outward-looking fashion?

Addressing these issues is one important focus of the research program of the International Trade Division of the World Bank. It has produced a number of methodological innovations in the traditional area of trade effects of RTAs and is now starting to tackle four new areas of research: the dynamics of regionalism (e.g., convergence, growth, investment, industrial location and migration), deep integration (standards, tax harmonization), regionalism and the rest of the world (including its effects on the multilateral trading system), and certain political economy dimensions of regionalism (e.g., credibility and the use of RTAs as tools of diplomacy).

In addition to thematic work, the program includes a number of studies of specific regional arrangements, conducted in collaboration with the Regional Vice Presidencies of the Bank. Several EU-Mediterranean Association Agreements have been studied and a joint program with the staff of the Latin American and Caribbean Region entitled “**Making the Most of Mercosur**” is under way. Future work is planned on African and Asian regional integration schemes.

Regionalism and Development findings have been and will, in future, be released in a number of outlets. Recent World Bank Policy Research Working Papers concerning these issues include:

Glenn Harrison, Tom Rutherford and David Tarr, “Economic Implications for Turkey of a Customs Union with the European Union,” (No. 1599).

Maurice Schiff, “Small is Beautiful, Preferential Trade Agreements and the Impact of Country Size, Market Share, Efficiency and Trade Policy,” (No. 1668).

L. Alan Winters, “Regionalism versus Multilateralism,” (No. 1687).

Planned future issues in this series include:

Eric Bond, “An Operational Model for Assessing Preferential Trading Arrangements”

Sherry Stephenson, “Standards, Conformity Assessments and Developing Countries”

Maurice Schiff and L. Alan Winters, “Regional Integration as Diplomacy”

Magnus Blomström and Ari Kokko, "How Foreign Investment Affects Host Countries" (WPS1745, March 1997)

Magnus Blomström and Ari Kokko, "Regional Integration and Foreign Direct Investment: A Conceptual Framework and Three Cases"

Anthony Venables and Diego Puga, "Trading Arrangements and Industrial Development"

Glenn Harrison, Thomas Rutherford and David Tarr, "Trade Policy Options for Chile: A Quantitative Evaluation"

In addition, **Making the Most of Mercosur** will be issuing papers over the next few months, including:

Alexander J. Yeats, "Does Mercosur's Trade Performance Raise Concerns About the Effects of Regional Trade Arrangements?" (WPS1729, February 1997))

Azita Amjadi and L. Alan Winters, "Transport Costs and 'Natural' Integration in Mercosur" (WPS1742, March 1997)

Claudio Frischtak, Danny M. Leipziger and John F. Normand, "Industrial Policy in Mercosur: Issues and Lessons"

Sam Laird (WTO), "Mercosur Trade Policy: Towards Greater Integration"

Margaret Miller and Jerry Caprio, "Empirical Evidence on the Role of Credit for SME Exports in Mercosur"

Malcom Rowat, "Competition Policy within Mercosur"

For copies of these papers or information about these programs contact Maurice Schiff, The World Bank, 1818 H Street NW, Washington, D.C. 20433.

An additional major outlet for World Bank-sponsored research on regionalism will be the Annual Bank Conference on Development in Latin America, 1997, Montevideo, June 30-July 2, 1997, organized by the Office of the Chief Economist and the Technical Department for Latin America and the Caribbean Region, with the support of the International Trade Division and the Economic Development Institute.

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1. Introduction

This paper attempts, for the first time we believe, to trace the effects of creating a regional trading arrangement on export and import prices. We commence by observing that it has long been known theoretically that preferential integration is likely to affect the terms of trade -- e.g., Mundell (1964) -- and by noting that the terms of trade are the most direct way in which regionalism affects non-partner countries. Indeed, the traditional method of assessing third country effects by examining whether the bloc imports more or less of the latter's exports is deeply flawed, and we view the present exercise as a step towards developing more satisfactory methods.

Section 2 sets out the background on the price effects of regional integration more fully. We then initiate an empirical test to determine, in one particular instance of integration, whether these effects can actually be identified. This begins in Section 3 which lays out a simple model of Bertrand competitors selling in a market which starts to offer one of them preferentially reduced tariffs. The prediction, not surprisingly, is that the preferred exporter will normally raise its pre-tariff price, while non-preferred one drops its, although this outcome is not completely inevitable. The model suggests that the effects of tariff changes are the same as those of changes in the costs of production and exchange rates in the exporting market. This is empirically useful because it allows us a far stronger empirical base for identifying the effects we postulate.

Section 4 gives details of the instance of regional integration we examine and the reasons for choosing it: we explore the effect of Spanish accession to the European Communities in 1986 on the prices of Spanish imports of finished manufactures from the United States, Japan, France, West Germany, Italy and the United Kingdom. Although

this is obviously only half of the terms of trade story, it seems an obvious place to start looking for price effects: the Spanish market is large enough to be worth competing for, differentiated products like finished manufactures offer the most likely locus for oligopolistic pricing, and the partners chosen are the main suppliers, at least in the engineering sectors. The exercise is conducted on the most detailed data available to us -- goods defined at the 5-digit level of the SITC(R).

Section 5 reports some results. After ruling out various 5-digit trade flows for data reasons or because of the presence of confounding policies, we have a sample of approximately 160 goods. The unit value series, which proxy import prices, are very noisy and the tariff aggregates rather crude. As a consequence it proves quite hard to extract information from the data. We concentrate on the relative prices of partner (i.e., EC) and non-partner sales to Spain, in the same way as the traditional assessments of the effects of regionalism are based on relative quantities or values of trade -- i.e., on import shares. Our estimates suggest that the relative tariff-inclusive prices of U.S. and Japanese sales in Spain relative to those of various EC suppliers respond to changes in relative costs and tariffs between member and non-member countries with an elasticity of 0.41-0.81 for the U.S. and 0.62-0.89 for Japan. This implies that, *ceteris paribus*, a tariff preference of, say, ten percentage points offered to EC suppliers causes the U.S. and Japan pre-tariff prices to fall relative to EC pre-tariff prices by approximately 2-6%, and 1-4% respectively. If this were manifest in a fall in the U.S. prices relative to the no-preference situation it would imply significant welfare losses for the United States on its sales to Spain. In fact, of course, at least part of the relative price change reflects increases in EC

supply prices, and to that extent the U.S. and Japan losses are relative only to a situation in which they also received the tariff reduction.

Identifying the terms of trade effects should be an important element in any economic assessment of regional trading arrangements, both for the distribution of gains among partner and for identifying the consequences on the rest of the world. Our results are, we believe, the first ever and are subject to a number of qualifications. However, they do suggest that such studies are feasible and that we can investigate the effects of regional integration on the rest of the world in a more direct and theoretically acceptable way than has been done heretofore.

2. Integration and the terms of trade

Perhaps the most central prediction of customs union theory is that entering a preferential trade arrangement alters a country's terms of trade. On the import side, this can occur in a simple Vinerian model through trade diversion; dearer supplies from partner countries (dearer in a national, pre-tariff, sense) displace cheaper supplies from non-partners and thus raise the average price of imports. In more sophisticated models with either increasing costs or differentiated products it occurs as increased demand for partner products, either created or diverted, raises their prices, while the opposite occurs for imports from non-partner. Figure 1 illustrates the simplest case. It refers to the market for an import supplied by two exporters, both subject to increasing cost (of supplying that market) and both subject to a tariff. When both face tariffs, total supply is the horizontal sum of the individual tariff-inclusive supply curves $S_N(t)$, for non-partners, and $S_P(t)$, for partners, viz., $S_{TOT}(MFN)$. When the partner is exempted from the tariff, the total supply

curve becomes the sum of $S_N(t)$ and S_P , viz., $S_{TOT}(CU)$, and the internal price falls from p_0 to p_1 . For the partner country, the price received rises from p_0' to p_1 for it faces no tariff in the custom union. For the non-partner, on the other hand, the price received falls from p_0' to p_1' , as demand for its tariff-ridden sales contracts.

On the export side, the story is more complicated but is still potentially important. Export prices to partner countries are likely to increase as exporters appropriate some or all of what was previously tariff revenue; those to non-partners might rise or fall depending on their substitutability in production for partner-bound exports; if integration has major effects on market structure -- for example, dramatically increasing levels of competition (Gasiorek, Smith and Venables, 1992) -- costs could fall sufficiently to bring all export prices down.

Clearly terms of trade effects are likely to figure significantly in the net welfare effects of a customs union or free trade area and also in their distribution across members. Perhaps more significantly, however, the terms of trade are the principal route through which integration schemes affect the rest of the world. We have argued in Winters (1996, 1997) that the usual approach to measuring this latter effect is deeply flawed: it entails asking whether non-members' shares of member countries' total imports have increased and equating increases with improvements in non-member welfare -- e.g., McMillan (1993) and IMF (1994). This is essentially a mercantilist calculus which, at face value, equates increases in welfare with giving away more goods! Of course, proponents of such a view argue, non-member exports are not given away but are sold for increased imports which raise non-member welfare directly. This is true, but as Winters (1997) shows, even in the simplest of models, the link from exports to welfare is indirect and not monotonic.

The two measures which do show a direct, necessary and sufficient connection to non-member welfare are non-member imports and non-member terms of trade,¹ and it is these that should be the focus of our measurement efforts.²

The effects of regional integration schemes on the terms of trade have not been entirely neglected. In an insightful, but largely unrecognized, article Mundell (1964) showed that, in a standard three-country model, if A and B offer each other mutual preferences, their joint terms of trade vis-à-vis the rest of the world (C) will improve. That is, the rest of the world will suffer. Petith (1977) uses Mundell's approach in a crude quantification of various instances of European integration, using essentially a very early and simple computable general equilibrium model. This is an *ex ante* study, drawing on no actual terms of trade data from the integration period. Similarly, Kreinin and Plummer (1992), assessing the effects of the European Communities' Second Enlargement and the formation of NAFTA, predict rather than measure the terms of trade effects on non-member countries.

More recent computable general equilibrium models have also generated predictions about the terms of trade. Most notably, in discussing Europe's "1992" exercise, Gasiorek, Smith and Venables (1992) identify the welfare effects for European countries of prospective changes in both export and import prices: interestingly they

¹ Anderson and Snape (1994) also observe the importance of terms of trade improvements.

² Winters (1997) also observes that, contrary to common usage, the discussion of non-member welfare has virtually no connection with the Kemp-Wan Theorem--Kemp and Wan (1976). Kemp and Wan discussed only the case in which the customs union chooses its external tariff to leave the prices and quantities of all non-members' trade unchanged and hence to leave non-members indifferent to the integration arrangements. They made no comment on the desirability or otherwise of any deviations from that position.

estimate quite significant declines in the EU's terms of trade (improvements in the rest of the world's terms of trade) as efficiency gains work through to export prices. Several papers in Francois and Shiells (1994) estimate the effects of NAFTA on the partners' terms of trade.

In the light of its theoretical importance it seems surprising that (as far as we know) there is no existing *ex post* study of the terms of trade effects, or even of the price effects, of regional integration. This paper attempts to start to put this situation right. It examines whether Spanish accession to the EC led to detectable changes in the prices of Spanish imports. This is obviously not sufficient to inform us about the terms of trade effects, but it seems a sensible place to start, given the theoretical results that exist.

3. A simple model

Although the notion that regional integration affects the prices in international trade requires no further justification, it is useful to examine more formally what form that relationship might take. We start with the simplest of models which might generate import price effects from a regional integration arrangement. We assume that the importing country (in our case below, Spain) imports a good from two sources and that it alters the tariff on one of those sources. The two sources supply differentiated versions of the good and we assume that this good is separable from all other expenditure in Spain and subject to two-stage budgeting. On the supply side we assume that only one firm from each country actually exports to Spain and that each has a cost function homogeneous of degree one in input prices; we assume that firms behave in a Bertrand fashion in the Spanish market and that this market is strategically separable from all others.

Thus while sales in other markets might affect the costs of supplying Spain, there is no attempt by firms to combine their interactions in different markets into a supergame. Each firm is assumed to maximize profits from Spanish sales in terms of its own currency.

The objective functions of the two exporting firms, member (which receives the tariff preference) and non-member, are as follows:³

$$Max_p \left\{ \frac{e}{\tau} p x(p, p^*, Q) - c(x, X) w \right\} \quad (1a)$$

$$Max_{p^*} \left\{ \frac{e^*}{\tau^*} p^* x^*(p, p^*, Q) - c^*(x^*, X^*) w^* \right\} \quad (1b)$$

where

- * the stars represent the member's variables, non-stars the non-member's;
- e is the exchange rate in terms of exporter's currency units per peseta;
- p is the price of sales in Spain including tariffs in peseta, we have also
 - a pre-tariff price \bar{p} such that $p = \bar{p} (1 + t)$;
- τ is the ad valorem tariff factor $(1+t)$, where t is the tariff on these exports;
- $x(p, p^*, Q)$ is Spanish demand for this exporter's products;
- $Q(Y, P, \bar{P})$ is the volume of imports of this good into Spain -- an aggregate
 - of the quantities supplied by the two exporters, x and x^* ;
- Y is Spanish total income;
- $P(p, p^*)$ is the price index for Spain's total imports of the target good -- it is
 - homogeneous of degree 1 in the individual prices;
- \bar{P} is an aggregate of all other Spanish prices;

³ The formalities of this model are quite close to those of Feenstra's (1989).

$c(x,X)$ is the marginal input/output ratio for the exporter -- the number of composite factor units required to deliver one unit of exports to Spain;

X is all other output by the exporting firm; and

w is the price of the composite factor.

The variables for the member country are defined analogously.

We treat all variables as non-stochastic and re-express the demand function in a form that combines both the partial price effect holding aggregate imports constant and the effect that arises as changes in individual prices affect the aggregate import price index, viz., $x = x(p, p^*, Y, \bar{P})$.

The first-order conditions are straightforward for non-member and member respectively,

$$m(p, p^*, Y, \bar{P}) \equiv p \left(1 + \frac{1}{\eta} \right) = z c_x \quad (2a)$$

$$m^*(p, p^*, Y, \bar{P}) \equiv p^* \left(1 + \frac{1}{\eta^*} \right) = z^* c_{x^*} \quad (2b)$$

where
$$z \equiv \frac{\pi w}{e} = \frac{(1+t)w}{e};$$

$\eta = p x_p / x$, the ordinary partial price elasticity of demand (also the appropriate “perceived” elasticity for Bertrand competitors); and $m(p, p^*, Y, \bar{P})$ is the marginal revenue function.

Equations (2a) and (2b) solve respectively to linear homogeneous functions

$$p = \pi(z, p^*, Y, \bar{P}) \quad (3a)$$

$$p^* = \pi^*(z^*, p, Y, \bar{P}) \quad (3b)$$

from which we calculate the effects on the export price of a change in the exporter's costs (w), exchange rate (e), tariff (t) or other output (X), and of a change in the competitor's price (including the effects of the tariffs on it).

Total differentiation of (2a) with respect to p, p*, z and (2b) with respect to p, and p* yields the following equation which must be solved to obtain the comparative statistics analysis of a shock to non-member country's costs, z.

$$\begin{pmatrix} \theta - \eta\gamma & \beta - \delta\gamma \\ \beta^* - \delta^*\gamma^* & \theta^* - \eta^*\gamma^* \end{pmatrix} \begin{pmatrix} \frac{dp}{dz} \frac{z}{p} \\ \frac{dp^*}{dz} \frac{z}{p^*} \end{pmatrix} = \begin{pmatrix} 1 \\ 0 \end{pmatrix} \quad (4)$$

where $\theta \equiv \frac{\partial m}{\partial p} \cdot \frac{p}{m}$, is the elasticity of marginal revenue with respect to price,

$\gamma \equiv \frac{dc_x}{dx} \cdot \frac{x}{c_x} \equiv \frac{c_{xx}x}{c_x}$ is the elasticity of marginal cost with respect to

the quantity of exports.

$\delta \equiv \frac{\partial x}{\partial p^*} \cdot \frac{p^*}{x}$ is the cross-elasticity of demand; and

$\beta \equiv \frac{\partial m}{\partial p^*} \cdot \frac{p^*}{m}$ is the elasticity of the target country's marginal revenue

with respect to the rival's price.

The sign of the elasticity of marginal revenue with respect to price, θ , depends only on the way in which the elasticity of demand evolves along the demand curve. Recalling that $m = p(1+\eta^{-1})$

$$\theta = \frac{\partial m}{\partial p} \cdot \frac{p}{m} = 1 - \frac{p^2 \eta_p}{m \eta^2} \quad (5)$$

$$> 1 \quad \text{if} \quad \eta_p < 0$$

$$< 1 \quad \text{if} \quad \eta_p > 0$$

Thus, for example, an isoelastic demand curve generates $\theta = 1$, while a linear demand curve, along which the absolute value of the elasticity rises with price ($\eta_p < 0$), gives $\theta > 1$. As long as the demand curve is not “too” convex, $\theta > 1$.⁴

The elasticity of marginal costs (γ) takes the sign of c_{xx} that is, it depends on whether marginal costs increase or decrease with scale. The cross-elasticity of demand δ , is assumed to be positive since the two varieties in question are taken to be substitutes.

One would normally expect $\beta < 0$; that is, holding own-price constant, an increase in the rival's price reduces the returns to selling one extra unit.

Solving equation (4) gives the following:

$$\frac{dp}{dz} \frac{z}{p} = \frac{(\theta^* - \eta^* \gamma^*)}{(\theta - \eta \gamma)(\theta^* - \eta^* \gamma^*) - (\beta - \delta \gamma)(\beta^* - \delta^* \gamma^*)} \quad (6a)$$

$$\frac{dp^*}{dz} \frac{z}{p^*} = \frac{(\delta^* \gamma^* - \beta^*)}{(\theta - \eta \gamma)(\theta^* - \eta^* \gamma^*) - (\beta - \delta \gamma)(\beta^* - \delta^* \gamma^*)} \quad (6b)$$

While equations (6) define the responses of prices to changes in tariffs in a fairly general context, the estimation of these responses requires more structure. Assuming that

⁴ Since $\eta_p = \frac{px_{pp}}{x} + \frac{x_p}{x} - \frac{px_p^2}{x^2}$, it requires the demand curve to be sufficiently convex for $\eta_p > 0$, which we will consider to be unlikely in our analysis.

the various elasticities are constant, we can use (3) to define (locally) the price functions (3a) and (3b) separately as:

$$\ln p = K + A \ln z + B \ln p^* + C \ln Y + D \ln \bar{P} \quad (7a)$$

$$\ln p^* = K^* + A^* \ln z^* + B^* \ln p + C^* \ln Y + D^* \ln \bar{P} \quad (7b)$$

Ideally one would estimate (7) directly, instrumenting $\ln Y$ and $\ln \bar{P}$ since the major shock to Spanish import prices -- accession to the EC -- probably affected aggregate prices and incomes as well. We chose not to do this initially for a number of reasons. First, it is useful to start with simpler and more direct methods in order to understand better the descriptive features of the data. Secondly, at least intuitively our actual method seems better able to isolate the effects we postulate. Third, it was not clear how best to instrument Y and \bar{P} . Fourth, it is obvious that system (7) should really have more equations since there are more sources of imports; deriving suitable cross-equation constraints in order not to overload our inadequate data is rather complex, and so is left for a later occasion.

We proceed, therefore, by considering a reduced form version of (7) and simplifying it by appealing to the assumptions made above. Separability and two-stage budgeting imply that the allocation of total demand for the target good across the two varieties x and x^* is homothetic. It depends only on relative prices, so that, in turn, p/p^* is independent of the aggregate demand or its determinants. Thus by focusing on the ratio of prices (difference in logs) we can eliminate the terms in Y and \bar{P} and, if they had been included, those in other varieties' prices. Our estimating equation is therefore

$$\ln p - \ln p^* = a + b \ln z - b^* \ln z^* \quad (8)$$

where the coefficients $a = \text{constant}$, b , and b^* are defined below.

To interpret equation (8), suppose that there is an increase in tariffs on non-members, i.e., $d\tau > 0$. Recalling that $p = \tau \bar{p}$ and that $z = \tau (w/e)$, (8) can be written in the form $\ln \bar{p} - \ln \bar{p}^* = (b - 1) \ln \tau - (b^* - 1) \ln \tau^* \dots$. Thus if $b = 1$ the change in the pre-tariff price of the member is exactly equal to that of the non-member. If $b < 1$ the member's pre-tariff price rise is smaller than that of the non-member, whereas if $b > 1$ then the opposite is true. We will show that all three cases are possible.

The final step is to define and interpret the coefficients b and b^* in terms of the elasticities above. Subtracting (6a) from (6b)

$$\begin{aligned} \frac{dp}{dz} \frac{z}{p} - \frac{dp^*}{dz} \frac{z}{p^*} &= \frac{d \ln \left(\frac{p}{p^*} \right)}{d \ln z} = \frac{(\theta^* - \eta^* \gamma^*) - (\delta^* \gamma^* - \beta^*)}{(\theta - \eta \gamma)(\theta^* - \eta^* \gamma^*) - (\delta \gamma - \beta)(\delta^* \gamma^* - \beta^*)} \\ &= \frac{1 - \left(\frac{\delta^* \gamma^* - \beta^*}{\theta^* - \eta^* \gamma^*} \right)}{1 - \left(\frac{\delta \gamma - \beta}{\theta - \eta \gamma} \right) \left(\frac{\delta^* \gamma^* - \beta^*}{\theta^* - \eta^* \gamma^*} \right)} \cdot \left(\frac{1}{\theta - \eta \gamma} \right) \end{aligned} \quad (9a)$$

We can also define the symmetric effect on the relative price given a shock to the costs of members,

$$\frac{d \ln \left(\frac{p}{p^*} \right)}{d \ln z^*} = \frac{(\delta \gamma - \beta) - (\theta - \eta \gamma)}{(\theta - \eta \gamma)(\theta^* - \eta^* \gamma^*) - (\delta \gamma - \beta)(\delta^* \gamma^* - \beta^*)} \quad (9b)$$

Examination of (9a) shows that there are two forces at work. It is easy to show that the second term is just the shift in price as non-member costs (tariffs) increase ignoring any interaction effects. The likely conditions that $\theta > 1$ and $\gamma > 0$ restrict this to lie between zero and one. The first term of (9a) captures the interaction between the two

firms. Figure 2(A) describes graphically what happens when an increase in the non-member's tariff shifts its reaction function by $(\theta - \eta\gamma)^{-1}$ from rf_1 to rf_2 , e.g., from point L to M. For convenience we define units such that $p = p^* = 1$, so that the ray from the origin to the initial equilibrium is the 45° line. The sign of the coefficient b is determined by whether the new equilibrium, N, lies above or below this ray. Stability of the Bertrand game requires that $\left(\frac{\delta\gamma - \beta}{\theta - \eta\gamma}\right)\left(\frac{\delta^*\gamma^* - \beta^*}{\theta^* - \eta^*\gamma^*}\right) < 1$ which we henceforth assume. If we then assume that both $\left(\frac{\delta\gamma - \beta}{\theta - \eta\gamma}\right)$ and $\left(\frac{\delta^*\gamma^* - \beta^*}{\theta^* - \eta^*\gamma^*}\right) < 1$, it is clear that $b < 1$: by the last assumption the interactive term is < 1 , and we have already argued that normally the shift term ≤ 1 . We will suppose that this is the normal case and we have represented it in Figure 2(A). Figure 2(B) shows that $b > 1$ is also possible even with stability if $\left(\frac{\delta\gamma - \beta}{\theta - \eta\gamma}\right) > 1$ (and if the shift is close enough to unity). Finally, figure 2(C) shows the case when $b < 0$, which is obtained with stability when $\left(\frac{\delta^*\gamma^* - \beta^*}{\theta^* - \eta^*\gamma^*}\right) > 1$.

We illustrate these results in a particularly simple case with constant costs ($c_{xx} = 0$), which implies that $\gamma = 0$, and isoelastic demand ($\eta_p = 0$), which implies that $\theta = 1$. Equation (9a) reduces to $(1 + \beta^*) / (1 - \beta\beta^*)$, whose sign depends only on the elasticity of marginal revenue with respect to the rival's price. In this case if the member country's marginal revenue were insensitive to the rival's price ($\beta^* = 0$) then the difference in price between non-member and member would change just by the shock in non-member costs. As the member country's marginal revenue becomes more sensitive to its rival's price

there would be greater relative losses for the non-member firm. If the elasticity of marginal revenue with respect to its rival's price is less than one and the stability requirement is satisfied, i.e., $\beta < 1$ and $\beta\beta^* < 1$. Then it is clear that the overall effect is less than one and therefore that a terms of trade loss occurs for non-members (Figure 2A). If $\beta > 1$ and $\beta\beta^* < 1$ then we get $b > 1$ (Figure 2B). If $\beta^* > 1$ and $\beta\beta^* < 1$, then $b < 0$ (Figure 2C).

Finally, the fact that the price functions (3) are linear homogeneous means that (8) should be homogeneous of degree 0, i.e., that $b = b^*$. We have experimented with this constraint and find that even though it is occasionally rejected statistically, imposing it never alters our conclusions materially. On data spanning more than a decade, equations that are not homogeneous are very uncomfortable, implying, as they do, the presence of money illusion.

4. The experiment and the data

One likely reason why no previous researcher has examined the price effects of regional integration is that doing so is empirically more complex than looking at trade flows. First, data on trade prices are generally absent, throwing one back on unit value series. These are very noisy, partly because prices frequently show significant variation across transactions, partly because intra-heading composition changes introduce spurious fluctuation and partly because of errors of measurement. As one moves to finer levels of the trade classification composition changes become less serious, but underlying price variation and measurement errors become worse because the reported data are averaged over fewer transactions. There is no doubt that one should resolve this trade-off by

aiming for the most detailed data available, but one should be correspondingly realistic about how much of the variance one can explain.

The second problem is that the price changes resulting from integration are likely to be small relative to those stemming from other causes. Thus while estimates that integration has affected trade volumes by 30-50% are not uncommon -- e.g., Aitken (1973) and Winters (1987) -- one does not expect to see price changes exceeding, say, 5-10%. Given all the other shocks that trade prices, integration shocks are likely to be well camouflaged in the price data. The combination of noisy data and small impacts implies that one is likely to need robust econometric methods to identify the effects we hypothesize.

We chose our experiment with the intention of maximizing the chances of identifying price effects, although in retrospect it is not clear that it was an ideal choice. We examine the effects of the Spanish accession to the European Communities in 1986 on the prices of Spanish imports of finished manufactures from major OECD sources. These sectors were selected as being the most likely to be subject to oligopolistic pricing -- price effects seem more likely to arise from imperfect competition than from upward-sloping supply curves because the new partner is rarely likely to be large enough to induce serious increases in costs for an exporter. These are also the markets for which our set of suppliers -- the United States, Japan, France, Germany, Italy and the United Kingdom -- are likely to be the predominant suppliers.

Spain appeared to be a good case study in that its accession is placed reasonably close to the center of our available sample period, 1980-1993, it is large enough that exporters would find its market worth competing for, and it did not belong to any free

trade arrangement prior to access that would confound the EC effect. The problem we face is that from the late 1970s, Spain offered EC suppliers tariff preferences on most goods, with discounts of 60% and, more commonly, 25% of the MFN tariff. Since the Spanish tariff very roughly averaged 20% before accession and the EC external tariff averaged roughly 5% after accession, the margin for EC suppliers often did not change very much.⁵ One solution to this is to extend the sample back into the 1970s to reflect the introduction of the preferences. However, difficulties over changes in quality (which are important in finished manufactures), changes in trade classification, small flows (which make unit values noisy) and the complications of the UK's position during its transition to full EC membership cause us to mistrust the earlier data somewhat.⁶

Spain's tariff was adjusted to EC levels over eight years following its accession in 1986. Thus for EC suppliers it converged from its 1986 starting point (including the EC discount) to zero in a series of roughly annual reductions of 10-15% of the starting tariff, while for non-members it converged on the EC common external tariff with the same time pattern, but with the steps being applied to the difference between the 1986 Spanish and EC tariffs. Thus even accession generated no dramatic steps in relative tariff levels which would allow a simple event analysis.

The most detailed unit value data for which reasonable time series are available are at the 5-digit level of the SITC(R) from the UN's COMTRADE. Corresponding data on tariffs were taken from *The International Customs Journal* published by the International

⁵ Prior to accession EU suppliers received a preference margin of $0.25 t_s$, where t_s is Spain's MFN tariff; after accession the margin is t_E , where t_E is the EU external tariff. If $t_s \approx 20\%$ and $t_E \approx 5\%$, which is roughly true of a significant share of headings, the preference margin hardly changes with accession.

⁶ Some results for 1970-1993 are given in the Appendix, however.

Customs Tariffs Bureau and the *Protocol to the General Agreement to Tariffs and Trade*, Geneva (1979) for years 1970-1986 and from information supplied by the Statistical Department of the Ministry of Finance of Spain for the years 1987-1994.⁷ Over the first period the tariff classification was organized according to the Brussels Tariff Nomenclature (BTN) and was matched to the unit value data using the UN's 4-digit BTN to 5-digit SITC(R) concordance. Within each 4-digit BTN category a simple average of the rates for individual tariff lines was taken. The Spanish tariff contains a "General" rate, a "GATT" rate reflecting bindings given under the GATT, and a discounted rate for EC suppliers. We took the lowest applicable of these for each heading and exporter. After 1987 the Spanish tariff was re-organized according to the Combined Nomenclature (CN); we matched it to our SITC(R) classification using a concordance provided privately by staff of UNCTAD, to whom we are most grateful. This change of basis may have redefined certain of our goods categories, and it certainly led to significant changes in the number of tariff headings contained in each category. Thus it may have introduced (a step) inaccuracy into some of our tariff series.

It proved necessary to drop certain 5-digit SITC(R) categories from the exercise. First, in a number of cases changes in the tariff classification significantly disturbed either the mapping of tariff data into SITC groups or the number of tariff headings averaged within a group. Second, Spain offered a concessionary tariff of 5% on imports of a list of capital goods on a year-by-year basis. Since we could not obtain the lists for every year

⁷ We are most grateful to Carlos Martinez Mongay and Vincente Isquierdo, for help in finding and managing the tariff data of the transition years. We are also grateful to Dan Gardener of the U.S. Department of Commerce for providing us with rare copies of *The International Customs Journals*.

and did not know the initial phase-in period for the concession, we dropped any 5-digit category which contained headings affected. Third, commodities were dropped from the exercise whenever they suffered more than four missing values over the main sample period of 1980-1993, and, finally commodities which were subject to quantitative restrictions in 1986 were also dropped. These excisions left us with samples of approximately 145 to 165 product categories, depending on the pairs of countries we examined.

Data on the wage bill and employment, from which wages have been derived, are taken at 2-digit level of the ISIC from the OECD's STAN database. Exchange rates were also taken from the OECD's STAN database. The indices of industrial production from which we derived the Capacity Utilization terms were taken at 3-digit level of the ISIC from the UN's UNIDO database.⁸ We treat all the variables mentioned in this paragraph as exogenous in our estimation, because they refer to substantially larger aggregations of goods than the 5-digit price data.

5. Results

The estimated model is:⁹

$$\ln \frac{p_{it}}{p_{it}^*} = a + b \cdot \ln \frac{(\tau w / e)_{it}}{(\tau w / e)_{it}^*} + c \cdot \ln \frac{CU_{it}}{CU_{it}^*} + \varepsilon_{it} , \quad (11)$$

⁸ Capacity Utilization is measured by the deviation of the logarithm of the appropriate 3-digit ISIC industry's index of industrial production from trend. Relative capacity utilization is the difference of two such variables.

⁹ Because costs of production might be sensitive to aggregate levels of output we add to the model, here and throughout, a term in relative pressure on capacity of the export industries of the two countries concerned.

where the subscript $i = (1, 2, 3, \dots, n)$ commodities, and $t = (1980, 81, 82, \dots, 1993)$ time. The RoW countries are represented by U.S. or Japan, and the EC (denoted by *) by one of the four major EC exporters to Spain, i.e., Germany, Italy, France, and the UK. We estimate equation (11) for every pair comprising one EC and one non-EC country.

Our first step was to estimate (11) on each of the products separately. That is effectively to index the parameters by i . Not surprisingly, with only 14 noisy observations, this produced a wide scatter of estimated coefficients and generated significant regressions for only 20-40 out of 145-165 commodities depending on which country pairs were used. The distributions of the estimates of b for each of the eight pairings of countries are given in Figure 3 and Figure 4. It is plain that while the estimates are concentrated in the correct region a priori -- between 0 and 1 -- their variances across commodities are huge. All this suggests is that the individual commodity estimations are not precise enough for us to be able to make useful inferences.

The weakness of the individual commodity results is disappointing and might, at first glance, be taken to indicate an incorrect model. However, it is very difficult to believe that prices are quite independent of costs, exchange rates and tariffs and insignificance indicates only the insufficiency of information to identify the postulated effects, not the irrelevance of the latter. Hence we proceed by combining the observations on the various commodities, first comparing results across regressions and then working with a cross-section of the time series.¹⁰

¹⁰ The ordering of the two concepts in this terminology is conscious.

Table 1 summarizes the estimates of b and the summary statistics for each country pair. The mean estimates of b are all statistically different from zero -- which would imply that post-tariff prices were independent of costs -- and statistically different from unity -- which would imply complete pass-through of costs -- in half the cases (U.S.-Italy, U.S.-UK, Japan-Germany, and Japan-France). The mean for the U.S.-UK sample, for example, is 0.65, which implies that, on average, a 1% increase in the difference in costs or tariffs between the U.S. and UK, i.e., in relative real wages or the difference in tariff between member and non-member countries, will cause tariff inclusive price differential to rise by 0.65%.

We now turn to the cross-section of time series. Combining the unit values for such different commodities obliges us to include separate constants for each product -- product fixed effects -- and to simplify the exercise we conduct the estimation on the deviations from individual means of the variables recorded in Section 3. We also correct for cross-sectional heteroscedasticity by deflating the data pertaining to each commodity by the estimated standard error from the individual estimate of (11) for that commodity. From here on, all pooled estimates will be corrected for heteroscedasticity in this manner.¹¹

Table 2 reports the estimates from the pooled samples for each of the eight country pairs. Due to the inherent noisiness of the unit value data only between 3% and 14% of the variance of the sample is explained by the equations, but it is apparent that the

¹¹ We have far too many commodities to make it feasible to include cross-commodity correlations in a SURE estimate. However, inspection of the matrix of cross correlations suggests that they are generally rather low.

regressions are all strongly significant statistically. Recall that all the explanatory power comes from the time-series dimension of the sample, the cross-commodity variation being absorbed by the product-specific constants. Thus, although the estimates from the individual regressions were widely scattered and the regressions themselves most frequently insignificant, there is enough information when we take them together to identify a set of effects common to all commodities.

Block A of table 2 shows that relative costs between member and non-member countries do have significant effects on relative tariff inclusive prices. To elaborate, it shows that the elasticities of relative prices with respect to relative costs between U.S.-UK, U.S.-Germany, U.S.-Italy, and U.S.-France are approximately 0.41, 0.73, 0.69, and 0.81, respectively: if relative U.S. costs increase by 10%, then the relative tariff-inclusive c.i.f. price of imports from the U.S. vis-à-vis the UK, Germany, Italy, and France in the Spanish market increases by 4.1%, 7.3%, 6.9%, and 8.1% respectively. This implies that part of any relative cost change is absorbed by exporters rather than passed on directly to consumers. We get similar results with Japan and its respective EC-4 members, with elasticities of 0.89, 0.68, 0.74, and 0.62 respectively -- 1% increase in J costs -- .11, .32, .26 and .38 decline in producer prices.

These elasticities also determine the effect of relative tariff changes on relative prices and thus on the extent to which producers absorb tariff changes. The coefficients are statistically significant from zero and, with one exception, significantly different from unity. The difference unity implies that, pre-tariff and post-tariff relative prices vary in opposite directions, a tariff reduction for an EC member country increasing its relative

pre-tariff export price. Alternatively stated, an EC tariff preference reduces the (relative) U.S. pre-tariff export price.

There is no evidence that Japan has a smaller pass-through to consumer prices than the U.S.; in fact the lowest coefficient is for the U.S.-UK pair. This result is perhaps surprising, given the alleged propensity of Japanese firms to “price to market.” On the other hand, the relative capacity utilization terms are positive and at least mildly significant for the U.S. pairs whereas they are quite insignificant for the Japanese bilateral cases. This suggests that while U.S. firms price exports according to domestic relative demand pressures, Japanese firms are less flexible.

The equations in block A of Table 2 represent our maintained hypothesis and they are consistent with the view that regional integration -- by affecting relative tariff burdens -- affects relative trade prices. That is, they suggest the conclusion that integration affects the terms of trade. We now seek to disaggregate the relative costs terms to identify separately the effect of changes in tariffs. Block B of Table 2 shows the results of tests for whether the tariff effect is significantly different from the combined effects of all costs, i.e., tariffs, wages and exchange rate. It reports estimates of the following model, where all variables are deviations from means:

$$\ln \frac{p_{it}}{p_{it}^*} = a + b_1 \cdot \ln \frac{\tau_{it}}{\tau_{it}^*} + b_2 \cdot \ln \frac{(w/e)_{it}}{(w/e)_{it}^*} + c \cdot \ln \frac{CU_{it}}{CU_{it}^*} + \varepsilon_{it} ,$$

as before $i = (1,2,3,...,n)$ commodities, and $t = (1980,81,82,...,1993)$ time.¹²

¹² This estimated model is no different from:

$$\ln \frac{\bar{p}_u}{\bar{p}_u} = a + (b_1 + b_2 - 1) \cdot \ln \frac{\tau_u}{\tau_u^*} + b_2 \cdot \ln \frac{(w/e)_u}{(w/e)_u^*} + c \cdot \ln \frac{CU_u}{CU_u^*} + \varepsilon_u$$

where \bar{p} is the pre-tariff price and $p_{it} = \bar{p}_{it} \tau_{it}$

None of the extra terms in tariffs is statistically significant -- indeed only one has a t-statistic above 1.2 -- so we can conclude that tariffs have no different effect on pricing than do other costs and that these effects are as reported in block A. Three of the eight tariff effects are negative -- meaning that the total effect of tariffs is less than that of costs and the exchange rate -- and five are positive. The latter tend to push the total tariff effects above unity, which implies that preferential reductions in the tariffs faced by the EC suppliers tended to be accompanied by reductions in their pre-tariff prices or increases in the United States' pre-tariff prices.

The most plausible explanation we have for these (insignificant) high tariff effects is that the tariff variables are proxying additional effects -- i.e., there are omitted variables that are highly correlated with our tariff data. One obvious candidate is that our sample covers a general deepening of integration within the European economy, so that, say, a tariff fall of 5% on European suppliers ($d \ln \tau^* \approx -0.05$) parallels a much stronger decline in overall trade frictions as, for example, the cost and complexity of customs procedures, transportation, foreign exchange dealings, product standards, and policy uncertainty declined to more normal levels. If these costs were previously borne by the European suppliers, what we measure as a tariff reduction would also include a significant reduction in other (unmeasured) costs. The latter would reduce member pre-tariff prices directly and, in our normal case, non-member pre-tariff prices would follow them down partially. Note that in this case, although the ratio of non-member to member pre-tariff prices would have risen as a result of the member's tariff preferences, non-member prices would still have fallen absolutely. Thus if we are right in associating declines in member tariffs with declines in European costs, the squeeze on non-member prices absolutely is reinforced not

mitigated. Let us re-iterate, however, statistically the tariff effects are not remotely significantly different from the general cost effects. In the results based on 1970-93 – see appendix table – the problem of “excessive” tariff effects are more marked. However, for the reasons noted above there are other reasons to worry about these results.

Recognizing that the unit value series are noisy and that we have a possible errors-in-variables problems, we also experimented with slightly more robust estimation techniques entailing data aggregation. Because we are exploiting only the time series variation, we need to maintain some variability in relative tariffs through time, so we re-estimated the equations on observations averaged over four three-year periods (1980-1982, 1983-1985, etc.) and one of two years (1992-1993).¹³ The resulting estimates, produced in Table 3, tell much the same story as the basic equations. All of the cost coefficients again are statistically different from both zero and unity although in the case of the Japan-UK country pair significantly greater than unity. Similarly, all the tariff effects in block B are insignificant and display a slight tendency towards being positive. The R-squares are, not surprisingly, a little higher, ranging now from about 6 %-41%. Overall, therefore, averaging the data to try to allow for noise and errors of measurement does not change our basic story.

So far we have treated the European suppliers as a separate and mutually competing, but it is also worth experimenting with treating them as a single unit. This helps to highlight the member/non-member dichotomy that lies at the heart of this work. To explore this question we aggregated all variables referring to the EC-4, weighting them

¹³ As in our previous results we have taken deviations from means and corrected for cross-sectional heteroscedasticity of all variables.

by their lagged shares in imports (in the commodity concerned) in the Spanish market.

The import shares were lagged by one period in order to avoid endogeneity problems.

Thus we estimate:

$$\begin{aligned} & \{\ln p_{it} - \sum_j \pi_{i,t-1}^{*j} \ln p_{it}^{*j}\} \\ &= a + b \cdot \left\{ \ln \left(\frac{pw}{e} \right)_{it} - \sum_j \pi_{i,t-1}^{*j} \ln \left(\frac{pw}{e} \right)_{it}^{*j} \right\} + c \cdot \{\ln CU_{it} - \sum_j \pi_{i,t-1}^{*j} \ln CU_{it}^{*j}\} + \varepsilon_{it} \end{aligned}$$

where π_{it}^{*j} is the individual EC-4 country's share of EC-4 exports of commodity i to Spain and $j = (\text{Italy, France, Germany, and UK})$. Table 4 shows the results of estimates using the same equations as reported in tables 2 and 3. The total costs coefficients are 0.77 for U.S.-EC-4 and 0.62 for Japan-EC-4, and again they are significantly different from both zero and unity. Similarly, when we tried to separate out the tariff effect we found we could not reject the hypothesis that the tariff effect was the same as the main costs effect for both Japan and the U.S.

6. Conclusions

Our objective in this paper has been to start to explore the terms of trade effects of regional economic integration. We have briefly shown why this is an appropriate measure of the welfare effects of integration, comparing it to the many *ex-post* studies which base their conclusions on changes in the import shares of member and non-member countries. We have demonstrated, through the use of a simple strategic model, how member countries might gain in their terms of trade, and non-members lose, through preferential

tariff lowering. And most importantly, we have shown that, though quite difficult, measuring such price effects of integration is feasible.

We believe that this is the first *ex-post* study of its kind and that it is a valuable addition to and an improvement over, previous *ex-post* studies on integration effects on the rest of the world. We have used finely disaggregated data on Spanish imports of finished manufactures from its major OECD trading partners, and despite their noisiness have found a consistent story over all of the country pairs examined. This shows that non-members, i.e., in our case the U.S. and Japan, suffered detectable terms of trade losses relative to EC competitors in Spanish import markets for differentiated goods.

In market structures where firms have some control over prices, it is not surprising that firms 'price-to-market' and that members reap terms of trade advantages with respect to non-members when their tariffs are preferentially lowered. Our present results do not allow us to identify how this relative change is apportioned between absolute increases in member prices and absolute falls in non-member prices, but the model suggests that the latter are very likely to some degree or other. In the present case the magnitudes involved are not very large. For our sample of finished manufactures the weighted average tariff preference for EC suppliers rose from 3.04% (10.18% - 7.14%) in 1986 to 5.20% (5.20 - 0%) in 1993. Converting to logarithms and using the elasticities from table 4(a) this translates into a relative price fall of 0.52% for the U.S. and 0.85% for Japan. Even on the most generous interpretation (i.e., that U.S. exporters experienced no price fall), this suggests that Japanese export prices fell by about 0.33%, and it is more likely that there was some fall in U.S. prices. If the latter were 0.10%, and if we apply the price falls to the

1995 volumes of trade in manufactures, the U.S. experiences losses of \$5.1 million and Japan losses of \$16.2 million, which are largely transferred to Spanish consumers.

The terms of trade costs to non-member countries of Spanish accession to the EU are small in the example above, partly because they suffered some discrimination before accession, partly because EC preferences are mostly rather small, and partly because the trade flows involved are relatively small. What the example suggests, however, is that elsewhere under less favorable circumstances the price-reducing effects of regional integration on non-members' exports could be significant.

Corresponding to the relative price reductions for non-European suppliers, are increases for the European ones -- and consequent losses to the Spanish exchequer. Using the same coefficients as previously suggests an increase in European suppliers' prices of 0.42%, which with 1995 manufactured import volumes from the EC-10 implies gains of \$232.6 million.

Despite the success of our empirical work, we recognize that this exercise is a partial one. There is ample space for further research, including on Spain's imports of other commodities and from other partners, changes in her export prices to Europe and the rest of the world, and the interaction between domestic and import prices. There is also ample opportunity to study the price effects of other examples of regional economic integration. Doing so will help us to gain a much more complete view of the effects of integration on both member and non-member nations.

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Figure 1

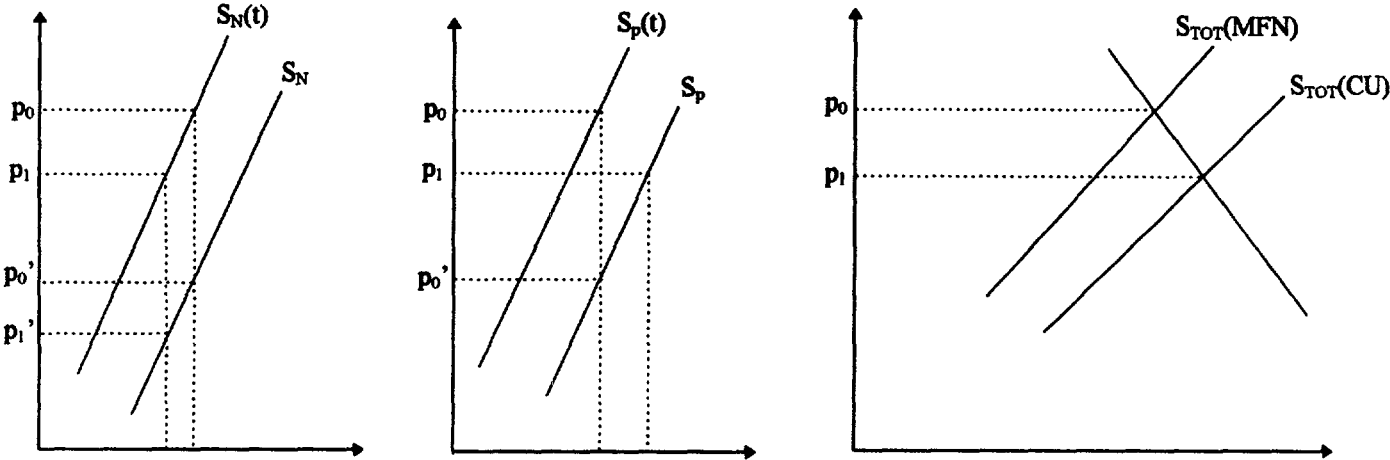


Figure 2

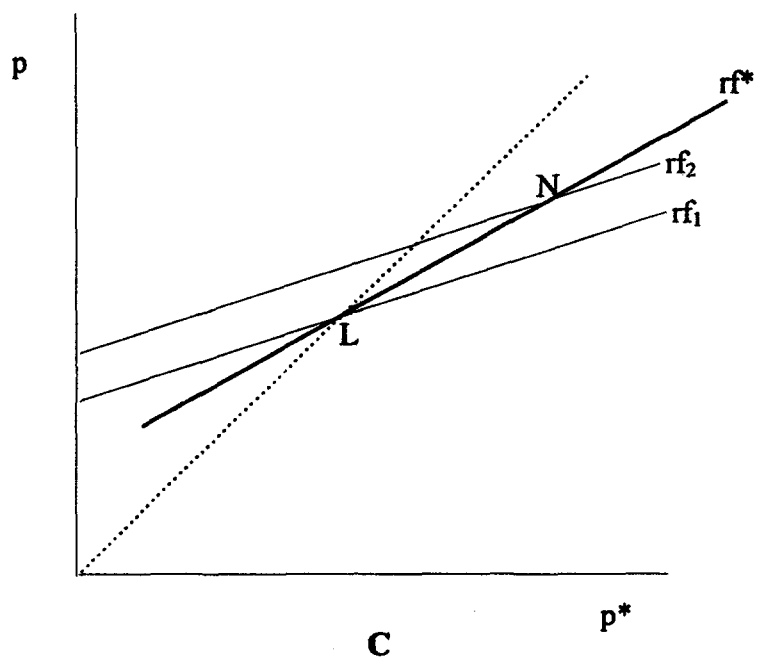
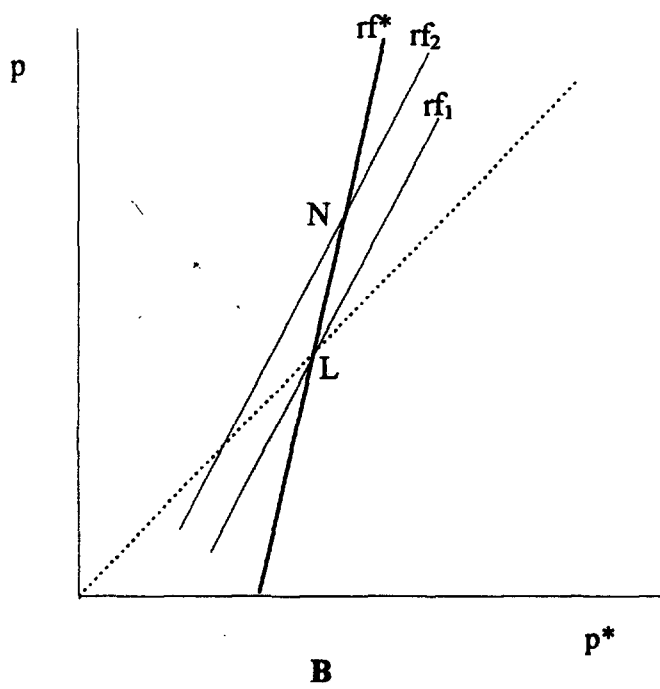
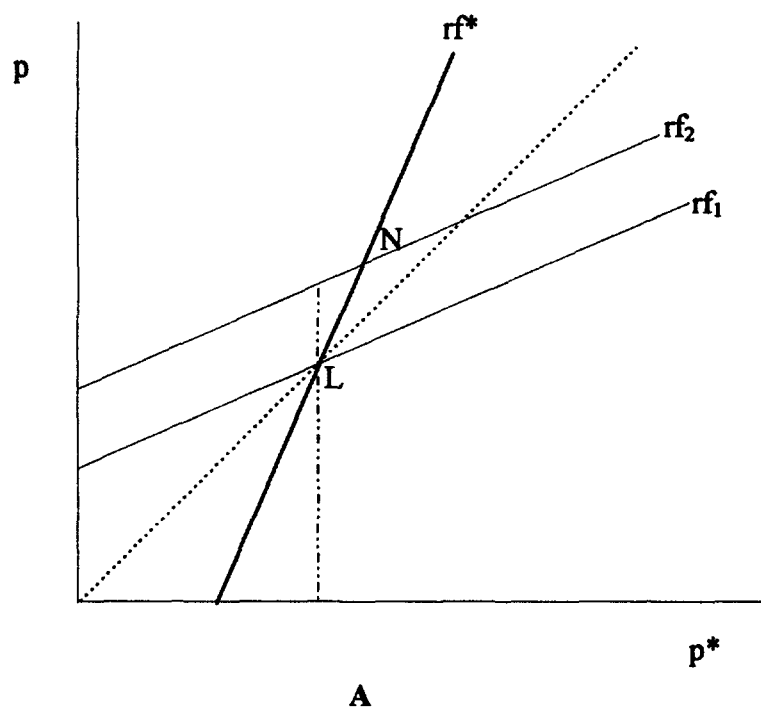


Figure 3: The distributions of the estimates of cost elasticity
USA-EC-4

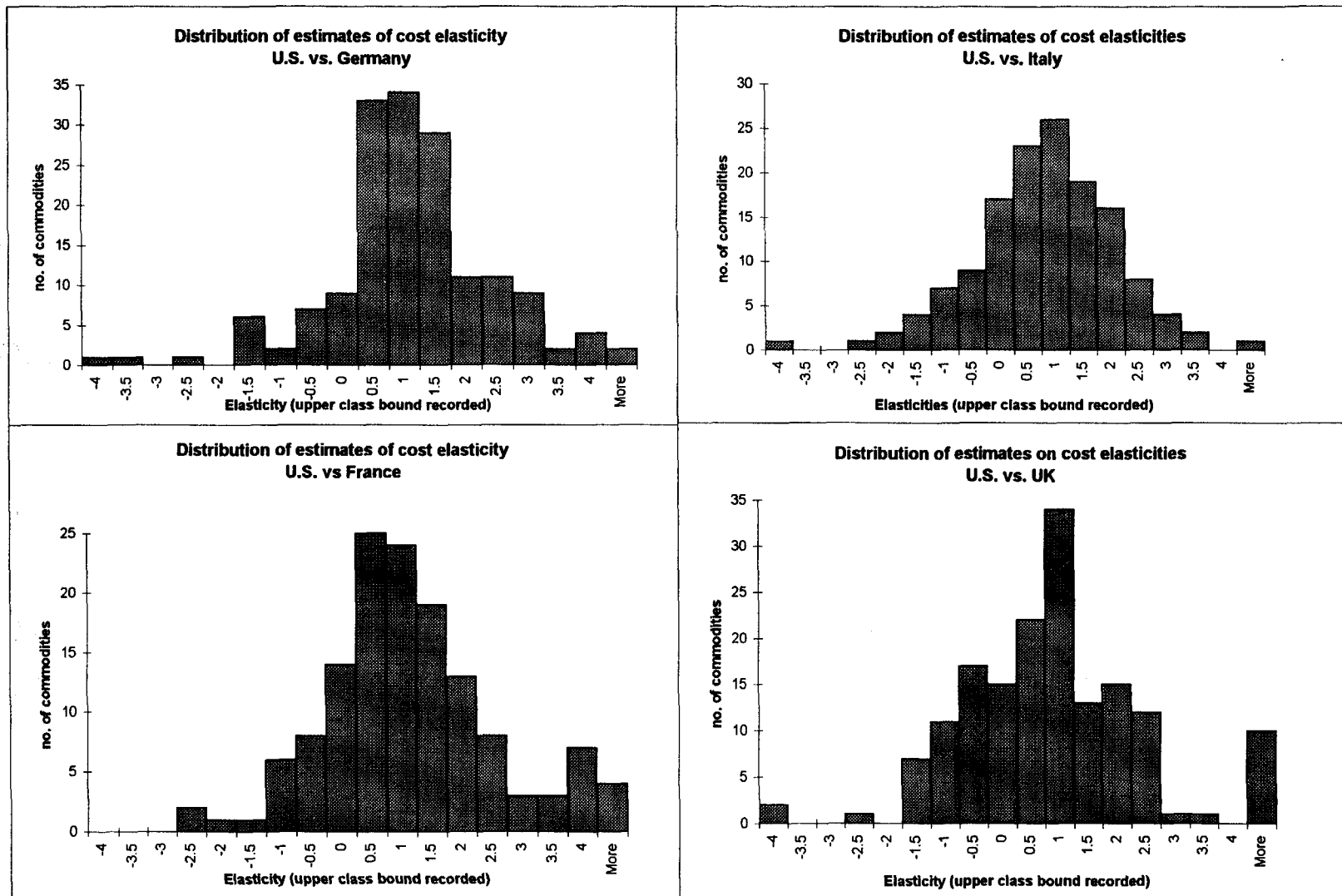


Figure 4: The distributions of the estimates of cost elasticity
Japan-EC-4

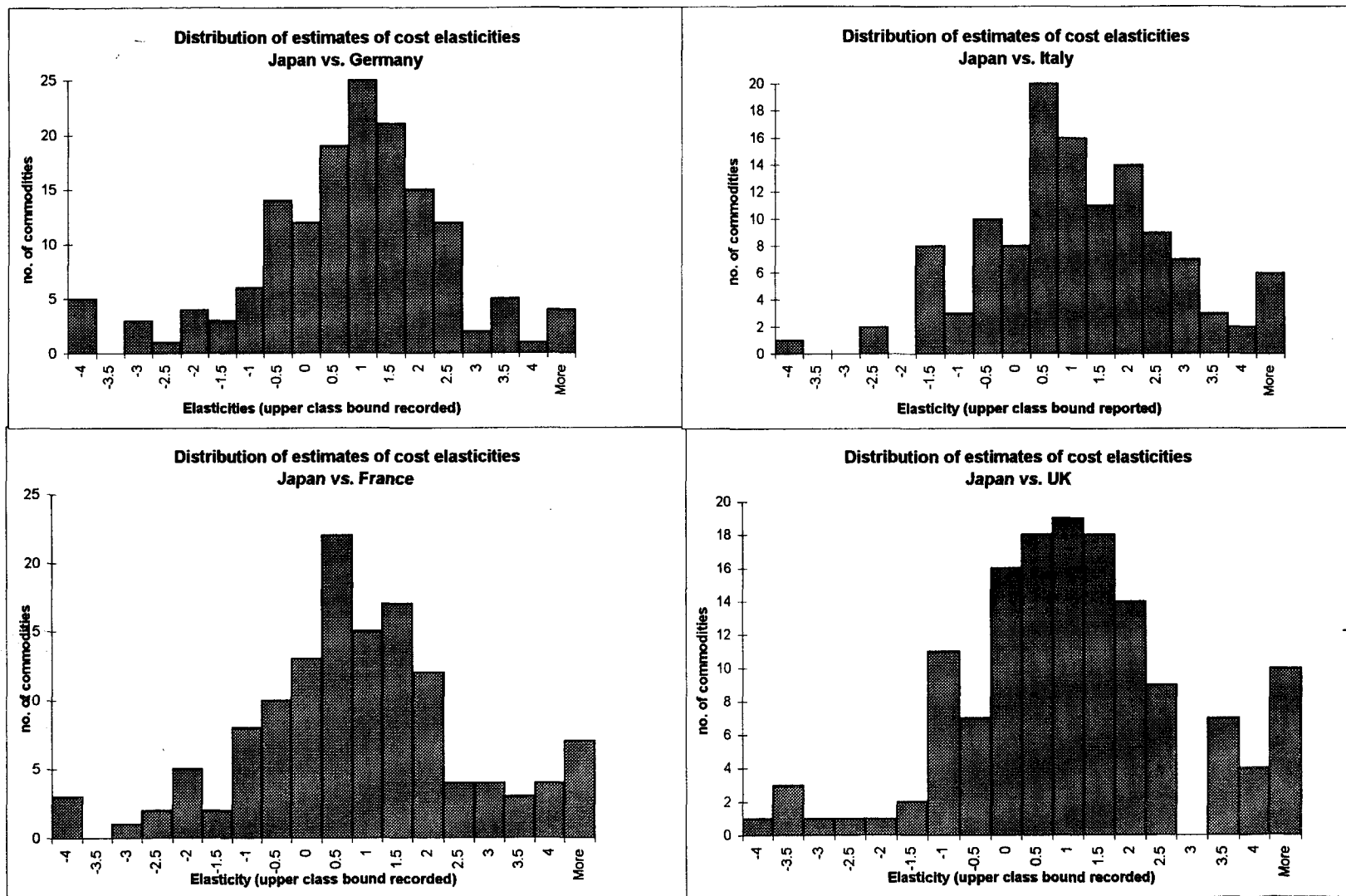


Table 1 Summary Statistics of Individual commodity estimates of equation (11)

	$\ln \frac{(\tau w / e)}{(\tau w / e)^*}$	$\ln \frac{CU}{CU^*}$	R-square	DW-Stat	F-signif**
<u>(1) U.S.-Germany</u>					
mean	0.8512	-0.1048	0.2812	1.5357	37/162
stdevp	1.3613	4.1393	0.2098	0.5954	
sd(mean)	0.1070	0.3252	0.0165	0.0468	
<u>(2) U.S.-Italy</u>					
mean	0.5698	0.5099	0.3028	2.0065	39/140
stdevp	1.3141	3.1873	0.2110	0.5686	
sd(mean)	0.1111	0.2694	0.0178	0.0481	
<u>(3) U.S.-France</u>					
mean	0.9221	-0.3023	0.2934	1.9563	36/138
stdevp	1.4583	4.2429	0.2135	0.6904	
sd(mean)	0.1241	0.3612	0.0182	0.0588	
<u>(4) U.S.-UK</u>					
mean	0.6450	-0.7249	0.2612	1.9496	37/161
stdevp	1.9468	4.1332	0.2087	0.5659	
sd(mean)	0.1534	0.3257	0.0164	0.0446	
<u>(5) Japan-Germany</u>					
mean	0.4607	-0.2175	0.2810	1.8265	39/151
stdevp	2.0654	4.1330	0.2085	0.5611	
sd(mean)	0.1675	0.3352	0.0169	0.0455	
<u>(6) Japan-Italy</u>					
mean	0.8647	-0.2804	0.2611	1.8127	20/120
stdevp	1.7417	3.4519	0.1857	0.6060	
sd(mean)	0.1590	0.3151	0.0170	0.0553	
<u>(7) Japan-France</u>					
mean	0.6130	-0.1992	0.2632	1.8960	28/132
stdevp	1.9980	3.7585	0.2205	0.6016	
sd(mean)	0.1739	0.3271	0.0192	0.0524	
<u>(8) Japan-UK</u>					
mean	0.8724	-0.0022	0.2995	1.8045	41/142
stdevp	1.8584	3.7194	0.2129	0.6118	
sd(mean)	0.1560	0.3121	0.0179	0.0513	

*Column of statistics are for coefficients of these variables.

**This column shows the number of regressions which are significant at 95% for each country pair examined.

Table 2 Parameter estimates from combined samples (1980-1993)*

(A)

	$\ln \frac{(\tau w / e)}{(\tau w / e)^*}$	$\ln \frac{CU}{CU^*}$	R-square	F-value
U.S.-UK	0.4084 (0.05147)	0.4787 (0.12353)	0.0328	37.4
U.S.-Germany	0.7302 (0.0385)	0.0554 (0.0806)	0.1399	181.2
U.S.-Italy	0.6927 (0.0402)	0.291 (0.0669)	0.1307	152.2
U.S.-France	0.8097 (0.0529)	0.1442 (0.0751)	0.1109	117.2
Japan-UK	0.8937 (0.0585)	0.2332 (0.1257)	0.1161	139.3
Japan-Germany	0.6798 (0.0571)	-0.0337 (0.0759)	0.0853	99.9
Japan-Italy	0.74773 (0.0644)	0.0571 (0.0489)	0.0765	80.5
Japan-France	0.62193 (0.0638)	-0.1595 (0.0743)	0.0588	56.1

(B)

	$\ln \frac{\tau}{\tau}$	$\ln \frac{(\tau w / e)}{(\tau w / e)^*}$	$\ln \frac{CU}{CU}$	R-square	F-value
U.S.-UK	0.1499 (0.8527)	0.4102 (0.0525)	0.486 (0.1238)	0.0328	24.9
U.S.-Germany	1.2258 (0.7788)	0.7423 (0.0392)	0.0611 (0.0833)	0.1408	121.7
U.S.-Italy	0.7967 (0.6968)	0.6948 (0.0402)	0.2856 (0.0671)	0.1313	101.9
U.S.-France	0.8817 (0.8)	0.8218 (0.0541)	0.1523 (0.0754)	0.1115	78.5
Japan-UK	-0.286 (0.799)	0.8944 (0.0586)	0.2375 (0.1263)	0.1162	92.9
Japan-Germany	0.4553 (0.633)	0.6778 (0.0571)	-0.0316 (0.076)	0.0855	66.7
Japan-Italy	-0.1394 (0.5807)	0.7471 (0.0645)	0.0581 (0.0491)	0.0765	53.7
Japan-France	-0.2515 (0.679)	0.6242 (0.0642)	-0.1625 (0.0748)	0.0589	37.4

*All estimates are corrected for cross-commodity heteroscedasticity and all variables are taken from deviations from means.

*Column of statistics are for coefficients of these variables. Parenthesis below are standard errors.

Table 3 Parameter estimates from combined samples (1980-1993)* Observations as three-year averages

(A)					(B)				
	$\ln \frac{(\tau w / e)^*}{(\tau w / e)^*}$	$\ln \frac{CU}{CU^*}$	R-square	F-value		$\ln \frac{\tau}{\tau^*}$	$\ln \frac{(\tau w / e)^*}{(\tau w / e)^*}$	$\ln \frac{CU}{CU^*}$	F-value
U.S.-UK	0.3814 (0.0596)	0.5185 (0.1376)	0.0565	24.5	U.S.-UK	0.0178 (0.8202)	0.3817 (0.061)	0.5184 (0.1379)	16.3
U.S.-Germany	0.6652 (0.0436)	-0.1637 (0.0792)	0.2495	136.8	U.S.-Germany	0.3505 (0.7109)	0.6679 (0.044)	-0.1643 (0.0792)	91.2
U.S.-Italy	0.5931 (0.0495)	0.2815 (0.0698)	0.2061	99.4	U.S.-Italy	-0.0337 (0.535)	0.593 (0.0436)	0.282 (0.0712)	66.2
U.S.-France	0.8567 (0.0594)	0.3878 (0.0722)	0.2231	104.1	U.S.-France	1.123 (0.7637)	0.8723 (0.0603)	0.3813 (0.07212)	70.2
Japan-UK	1.138 (0.0603)	0.2062 (0.0857)	0.3477	215.9	Japan-UK	0.1775 (0.8152)	1.139 (0.0604)	0.203 (0.0869)	143.7
Japan-Germany	0.6622 (0.066)	0.0198 (0.0745)	0.1575	76.1	Japan-Germany	0.5709 (0.6413)	0.6599 (0.06609)	0.0217 (0.0746)	50.9
Japan-Italy	0.8612 (0.0544)	0.1947 (0.0354)	0.2737	142.8	Japan-Italy	-0.6013 (0.5072)	0.8593 (0.0544)	0.2012 (0.03583)	95.7
Japan-France	0.5168 (0.0237)	0.0833 (0.0725)	0.4138	253.8	Japan-France	0.00858 (0.7061)	0.5165 (0.0368)	0.0835 (0.0742)	168.9

*All estimates are corrected for cross-commodity heteroscedasticity and all variables are taken from deviations from means.

*Column of statistics are for coefficients of these variables. Parenthesis below are standard errors.

Table 4 Parameter estimates from combined samples (1980-1993)*
U.S. and Japan vs. weighted average of EC-4

(A)

	$\ln \frac{(\tau w / e)}{(\tau w / e)^*}$	$\ln \frac{CU}{CU^*}$	R-square	F-value
U.S.-EC4	0.771 (0.0399)	0.286 (0.0956)	0.1825	187.8
Japan-EC4	0.625 (0.0582)	-0.138 (0.0681)	0.0728	64.5

(B)

	$\ln \frac{\tau}{\tau^*}$	$\ln \frac{(\tau w / e)}{(\tau w / e)^*}$	$\ln \frac{CU}{CU^*}$	R-square	F-value
U.S.-EC4	0.723 (0.4708)	0.772 (0.0399)	0.307 (0.0965)	0.1836	126.1
Japan-EC4	-0.266 (0.376)	0.623 (0.0583)	-0.1412 (0.0682)	0.073	43.2

Table 5 Parameter estimates from combined samples (1980-1993)
Observations as three-year averages*
U.S. and Japan vs. weighted average of EC-4

(A)

	$\ln \frac{(\tau w / e)}{(\tau w / e)^*}$	$\ln \frac{CU}{CU^*}$	R-square	F-value
U.S.-EC4	0.676 (0.0528)	0.327 (0.102)	0.2084	86.6
Japan-EC4	0.608 (0.0885)	-0.0458 (0.0926)	0.0809	29.1

(B)

	$\ln \frac{\tau}{\tau^*}$	$\ln \frac{(\tau w / e)}{(\tau w / e)^*}$	$\ln \frac{CU}{CU^*}$	R-square	F-value
U.S.-EC4	0.677 (0.548)	0.678 (0.0528)	0.343 (0.103)	0.2103	58.3
Japan-EC4	0.174 (0.496)	0.609 (0.0886)	-0.044 (0.0928)	0.0811	19.4

*All estimates are corrected for cross-commodity heteroscedasticity and all variables are taken from deviations from means.

*Column of statistics of the coefficients of these variables. Parenthesis below are standard errors.

Appendix Parameter estimates from combined samples (1970-1993)*

(A)

	$\ln \frac{(\tau w / e)}{(\tau w / e)^*}$	$\ln \frac{CU}{CU^*}$	R-square	F-value
U.S.-UK	0.3517 (0.0428)	0.2986 (0.0564)	0.0176	33.8
U.S.-Germany	0.555 (0.0385)	0.4856 (0.0563)	0.0639	130.4
U.S.-Italy	0.2474 (0.0339)	0.237 (0.051)	0.0158	28.6
U.S.-France	0.391 (0.0386)	0.226 (0.0368)	0.0327	58.3
Japan-UK	0.624 (0.0471)	0.133 (0.052)	0.0751	142.3
Japan-Germany	0.6951 (0.0383)	0.015 (0.0286)	0.1487	309.2
Japan-Italy	0.981 (0.0403)	-0.00026 (0.0244)	0.185	375.3
Japan-France	0.674 (0.0416)	-0.0195 (0.0292)	0.0938	163.3

(B)

	$\ln \frac{\tau}{\tau}$	$\ln \frac{(\tau w / e)}{(\tau w / e)^*}$	$\ln \frac{CU}{CU}$	R-square	F-value
U.S.-UK	-0.635 (0.516)	0.335 (0.0449)	0.317 (0.0584)	0.018	23.1
U.S.-Germany	3.382 (0.4754)	0.682 (0.0387)	0.362 (0.0586)	0.076	104.9
U.S.-Italy	4.518 (0.432)	0.408 (0.0368)	0.181 (0.0502)	0.0451	56.0
U.S.-France	3.178 (0.499)	0.555 (0.0462)	0.2202 (0.0366)	0.0439	52.8
Japan-UK	-1.074 (0.573)	0.67 (0.0531)	0.165 (0.0577)	0.076	96.1
Japan-Germany	0.286 (0.438)	0.683 (0.0425)	0.0137 (0.0287)	0.149	206.2
Japan-Italy	2.0561 (0.5182)	0.844 (0.053)	0.00026 (0.0244)	0.1887	256.6
Japan-France	2.162 (0.5347)	0.572 (0.0485)	-0.016 (0.0292)	0.0984	114.8

*All estimates are corrected for cross-commodity heteroscedasticity and all variables are taken from deviations from means.

*Column of statistics are for coefficients of these variables. Parenthesis below are standard errors.

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